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Education and Economic Growth in Germany before the Second World War. An Econometric Analysis of Dynamic Relations

Claude Diebolt / Javier Litago*

»Economic growth is a long-term process whose features can properly be observed only in a historical perspective."

[KUZNETS, 1969, p. 191]

Abstract: The nature of the dynamic relations between education and economic growth is far from having been determined by economists or historians and is probably one of the longest unsettled controversies in the field of economic history. In the context of this controversy the paper is devoted to the dynamic relations between public expenditure on education and national income in Germany before the Second World War.

Introduction

The relationship between education and economic growth has polarized opinions. The question of whether education *favours* or *hinders* growth has been asked and examined in static and dynamic terms by different writers [Tortella, 1990]. Some authors have resorted to a kind of cost-benefit analysis, comparing the losses and gains derived from education [Becker, 1964]. Here, the so-called static arguments are deliberately disregarded in favour of the dynamic consequences of education because the static consideration of gains (or losses) is irrelevant to the appraisal of the implications of education for economic growth.

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In this respect, the results of empirical investigations concerning the relationship between education (public expenditure) and economic growth (national income) in Germany before the Second World War are presented here. Rather than venture into new fields of the theory of growth [Barro and Sala-i-Martin, 1995], we confine ourselves to presenting certain new data and the results of several tests of existing theory. We have tried to cover somewhat long periods as far as this was possible with the statistical data available. Many past analyses of growth suffer from the fact that the time-span investigated was only one or two decades. Here, we have tried to use statistical data starting in the middle of the nineteenth century.

I. Methods

The use of mathematical methods in social history is in its infancy. It is not surprising therefore that advocates of the quantitative approach should find their methods subjected to criticism. This is especially true in the case of the application of stochastic processes[1] Some of this criticism doubtless arises from reluctance to change established methods and habits of thought, but some deserves serious consideration. Our objective here is not to contribute to the debate but to clarify the nature and especially the limitations of our methods. The complexity of social phenomena is frequently highlighted. This is expressed by saying that social situations are far too complicated to allow mathematical study and that to ignore this fact is to be led into dangerous oversimplification. The premise of this objection must be accepted.

Social phenomena are exceedingly complex and our models are bound to be simplifications. Even if it is allowed that our model need only reproduce the relevant features of the real process, the problem remains formidable. However, we would argue that there is no alternative to simplification. The basic limiting factor is not the mathematical tools available but the ability of the human mind to grasp a complex situation. There is no point in building models whose

[1] A stochastic process is one which develops in time according to probabilistic laws. This means that we cannot predict its future behaviour with certainty. The most that we can do is to attach probabilities to the various possible future states. Stochastic models of social phenomena have been constructed in the past with different objects in view. For our purposes it is useful to distinguish one main function of models: that of giving insight into and understanding the phenomenon in question. The investigation begins with the collection of data on the process and the formulation of a model which embodies the observed features of the system. This we shall describe as the model-building stage. The next step is to use the model to make predictions about the system which can be tested by observation. This activity will require the use of mathematical reasoning to make deductions from the model and will be referred to as model-solving. The final step is to compare the deductions with the real world and to modify the original model if it proves to be inadequate. This is the procedure of model-testing.

ramifications are beyond our comprehension. Perhaps the only safeguard against oversimplification is to use a battery of models instead of a single one. Any particular model will be a special case of the more complex model needed to achieve complete realism. Greater confidence can be placed in conclusions that are common to several special cases than in those applicable to a single arbitrarily selected model. However, in a new and rapidly developing subject, it would be premature to claim that this were a complete and systematic treatment.

II. Data

The basic procedure involves the assembly of long time-series of data (from 1850 to 1937) on both public expenditure on education and national income. The choice of starting date was guided by several considerations. 1850 was selected partly because of the difficulties in estimating capital formation and industrial growth before that date. Hoffmann's classic study of German economic growth starts in 1850 (Figure 1) and the difficulties involved in creating time-series for earlier years would have been daunting [Hoffmann *et* al., 1965]. Diebolt's study on the long-term evolution of public expenditure on education in Germany starts in 1829 [Diebolt, 1994]. In this paper, the data drawn up using quantitative history methods show the outlay devoted to primary, secondary and higher education (Figure 2) with 1937 representing the last normal year before the Second World War.

The figures from the First World War onwards in current marks obviously gives a very distorted picture of real growth. They must therefore be weighted using prices in order to approach the volume of expenditure. However, weighting of prices over a long period raises many methodological and theoretical problems. The problem of the formulation of the indexes used and especially the linking of successive indexes with different weighting systems is not discussed here. It is simply reminded that a picture of price changes over a long period is a mental exercise and one does not know to what extent it reflects the variations of the measuring instrument or real variations in productivity.

The picture of prices resulting from a juxtaposition of indexes - not permitted in theory - doubtless gives an approximation of real movements over relatively short periods but can in no way be used to compare prices on two dates that are distant from each other. In short, whatever the index used, it will always be the target of justified criticism as an index can only reflect certain aspects of economic reality. However, such calculations are essential even if they must be performed with retail or wholesale price indexes, which are poorly suited to the objective. A compromise was finally decided with use of two-thirds weighting by retail prices and one third by wholesale prices.

Figure 1

National Income in Germany, 1850-1937

Constant marks

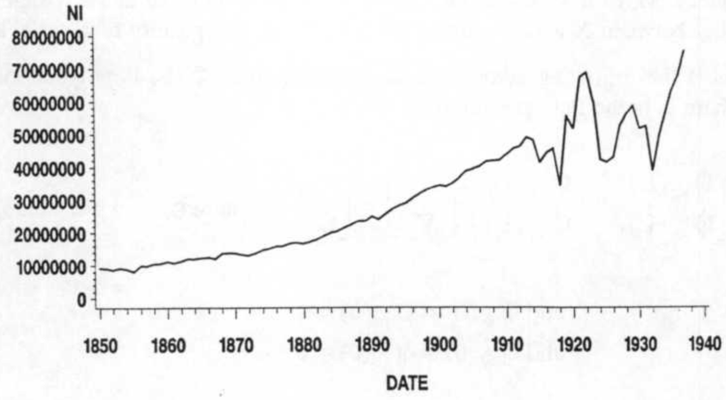
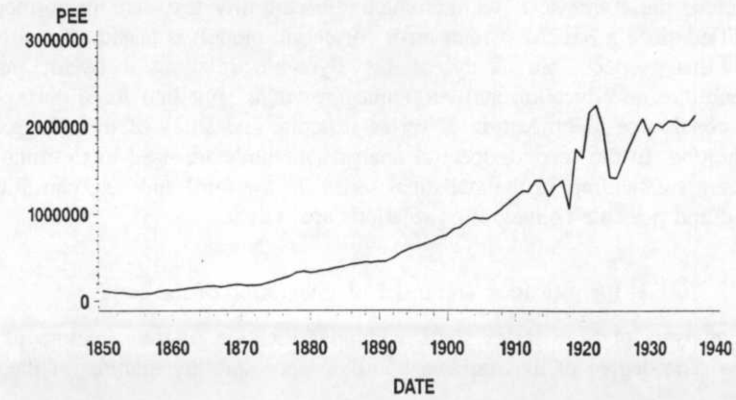


Figure 2

Public Expenditure on Education in Germany, 1850-1937

Constant marks



III. Empirical Findings

Causality as used by Granger refers to the anteriority of a phenomenon in relation to another [Granger, 1969]. More precisely, it is said that X causes Y if past values of X contain information that are not contained in the past of Y but which significantly improve its forecasting. This reasoning is performed within the framework of a VAR (vector auto regression) model that can envisage all causality between X and Y without prejudicing the exogeneity of a variable.

This is the following autoregressive system (where $\Phi_{-}(L)$ is polynomial in L), where L is the lag operator such that $L^j Z_t = Z_{t-j}$:

$$\begin{bmatrix} \phi_{xx}(L) & \phi_{xy}(L) \\ \phi_{yx}(L) & \phi_{yy}(L) \end{bmatrix} \begin{bmatrix} X_t \\ Y_t \end{bmatrix} = \begin{bmatrix} \varepsilon_{xt} \\ \varepsilon_{yt} \end{bmatrix}, \text{ where } \varepsilon_t, \text{ is white noise,}$$

$$\text{and } \phi_{xx}(0) = \phi_{yy}(0) = 1$$
$$\text{and } \phi_{xy}(0) = \phi_{yx}(0) = 0$$

It is said that X does not cause Y (and respectively Y does not cause X), if $\Phi_{yx}(L) = 0$ (respectively $\Phi_{xy}(L) = 0$).

The use of this type of modelling to test causality is only correct in certain hypotheses. First of all, the first and second order moments in the series must be finite. In other words, one must work with stationary data. The presentation of our results therefore starts with the identification of the structure of the statistical series and study of their degree of integration in order to reason according to their first differences if the series are integrated of order 1, that is to say non-stationary in level but stationary after differentiation. The existence of a cointegration relation indicates causality in at least one direction. In such a situation, the framework within which the causality tests are performed is modified since a VECM (vector error correction model) is tested.

In this respect, our study of the dynamic relations between public expenditure on education and economic growth is split into three parts. The first consists of identification of series structure and study of their degree of integration. In the second, spectral analysis methods are used to examine the inherent fluctuations in the statistical series. In the third and last, causality is tested and possible cointegration relations are sought.

1. Identification and order of integration of the series

Our analysis of time series starts with identification of the structure of the series. The degree of the stationary and variance stability features of the two variables chose were studied. Examination of the graphs of the original series (Figures 1 and 2) shows the need for logarithmic transformation in order to achieve variance stability. The estimation and representation of the autocorrelation and partial autocorrelation (AF and PAF) functions and the spectral density function enabled us to detect two major features.

For the national income (NI), AF displayed slow decrease to lags 28-30. This is characteristic behaviour in non-stationary processes. A very significant peak representing an autoregressive process of order 1 was observed in the first lag of the PAF. A very slight peak was observed in lag 6. In addition, the spectral density of this variable was concentrated at low frequencies (0), i.e. in an infinite period, thus confirming its non-stationarity.

With regard to public expenditure on education (PEE), an exponential decrease in the AF was observed as far as lags 27-29. Here, the PAF displayed a very significant peak in the first lag that is characteristic of the non-stationarity of this series that, as a preliminary approximation, displays the behaviour of an AR1 process. In addition, the strong spectral density at zero frequency confirms the non-stationarity of this series. A closer study of the stationarity conditions followed this first stage of identification of the structure of the variables. Various tests of unit root were used for this with the main aim of determining the order of integration of the two series.

Tests to detect the presence of unit roots in the generating process of a series (and whose presence implies the non-stationarity of the series) are the type most commonly employed to examine the order of integration of series. Wetherefore drew up a regression of the first differences of each variable on a constant, on a trend, adding p lags to the differences themselves as follows:

$$\nabla x_t = \alpha + \beta t + \gamma x_{t-1} + \sum_{j=1}^{k-1} \delta_j \nabla x_{t-j} + \varepsilon_t \quad \text{where p} = 1, 2, \ldots$$

This was to test the null hypothesis of the presence of a unit root a = 0 against a < 0 using the augmented Dickey-Fuller tests (ADF) with a (DF) constant, constructed like a Student's t test (the relation of a coefficient to its standard deviation), but whose distribution is not that of a Student test. The DFT statistic (with constant and trend) is constructed using the same principle but with time added to the initial regression as an explanatory variable [Dickey and Fuller, 1979, 1981].

After the test for the presence of a unit root, we sought the presence of a second root, that is to say we verified the hypothesis that the process of the LNI and LPEE series were integrated of order 1,I(1) and not integrated of order 2, I(2). Once the existence of strong residual autocorrelation had been verified in the autoregressions of the two variables, we applied the augmented version of the "Weighted Symmetric" (WS) estimator (Pantula *et al.,* 1994), Dickey-Fuller, ADF and Phillips-Perron[2] (PP) tests (Table 1).

The order of autoregression in the ADF and WS tests was determined using Akaike's information criterion rule plus 2, (AIC+2) for both the variables studied (Pantula *et al.*, 1994). The optimum order of the corresponding ADF test was used for the PP tests. The upper limit was set at 10 lags (approximately 10% of the number of observations). The results of these tests according to the first rule are set out in Table 1 (upper part). We then used as the second rule the residual autocorrelation criterion observed using the following tests: Durbin's h test, the Breusch-Godfrey LM test and the Ljung-Box Q(25) test. As the true order (p) of the processes in each series was unknown, we used Table 8 of Pantula *et al.* (1994, p. 457) to obtain the critical values of the tests for an order $\hat{p}$ estimated using the following criterion: min ($\hat{p}$ AIC+2, 10).

For the LNI series (first differences: ∇), this criterion selected an order of 9 (WS test) and an order of 7 (ADF and PP tests). The statistics of the WS test, $\hat{\tau}_{ws,\tau}$, was –2.17 (Table 1); this was less than the critical value at a 5% significance level, that is to say –3.36 (n=87<100 observations), with 0.53 asymptotic probability. This test did not reject the hypothesis of the present of a unit root. In addition, the statistic $\hat{\tau}_{\tau}$, of the ADF test was –0.73, less than the critical value –3.45, with asymptotic probability of 0.97 (P value). The null hypothesis at a 5% significance level in the presence on a unit root was thus accepted. The PP test confirms this hypothesis with lower probability (4%).

Table 1. Results of the Weighted Symmetric (WS), Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests on the LNI and LPEE series.

TESTS	SERIES: ∇ LNI	SERIES: ∇ LPEE
AIC+2	WS: -2.0, k=9; ADF: -2.0, k=7	WS: -1.81, k=2; ADF: -1.9, k=7
ADF	(k=7) -0.73 P value = 0. 97	(k=7) -0.60 P value = 0.98
WS	(k=9) -2.17 P value = 0.53	(k=2) -2.74 P value = 0.17
PP	(k=7) -22.29 P value = 0.04	(k=7) -17.71 P value = 0.11
TESTS	SERIES: ∇^2 LNI	SERIES: ∇^2 LPEE
AIC+2	WS: -2.1, k=6; ADF: -2.0, k=6	WS: -1.80, k=6; ADF: -1.9, k=6
ADF	(k=6) -4.64 P value = 0.001	(k=6) -5.65 P value = 0.00001
WS	(k=6) -4.40 P value = 0.00088	(k=6) -3.89 P value = 0.006
PP	(k=6) -76.43 P value = 1.9xd-07	(k=6) -72.19 P value = 0.0000005

For the LPEE series, the AIC+2 criterion selects an order of 2 for the WS test and an order of 7 for the ADF and PP tests. The statistic $\hat{\tau}_{ws,\tau}$ has a value of –2.74, below the 5% critical level, i.e. –3.36 (n=87<100 observations) with probability of 0.17 (P value). This test did not reject the null hypothesis of the presence of a unit root. In addition, the ADF test estimation gives statistics $\hat{\tau}_{\tau}$ a value of –0.60, less than the 5% critical level, i.e. –3.45, with probability of 0.98. This result supports the hypothesis of a process I(1) for the description of the LPEE series. The PP test confirms the presence of a unit root, with a P value of 0.11.

[2] Cf. Dickey-Fuller (1979, 1981), Phillips, P. (1987), Phillips and Perron (1988), and Pantula *et al.* (1994).

As an extension of this, the following conclusions are reached when the residual autocorrelation in the selection of order AR is considered for the models of the two variables studied. Regression must be increased to the fifth lag for the ADF test on the LNI series. With this, h changed from –4.66 to 0.6 between the fourth and the fifth lag. Here again, the hypothesis of absence of autocorrelation in the residuals is accepted. In the LPEE series, statistic h changes from –2.0 to 1.0 when the number of lags in the ADF autoregression is increased from 4 to 5, making it possible to reject the alternative autocorrelation hypothesis. This was also confirmed by the Breusch-Godfrey LM test and the Ljung-Box Q(25) test.

ADF tests were then performed on regressions of order 5 for the two variables studied. The statistic $\hat{t}_\tau$ was estimated to be –0.50 for the LNI series, that is to say lower in absolute value than the 5% critical threshold of –3.45 (Table 8.5.2. Fuller, 1976). The value was also distinctly lower than the 5%, –3.4652, and 10%, –3.2519, thresholds of McKinnon (1991). In the light of these tests, we assume the null hypothesis of the presence of a unit root in this series. In addition, on the basis of the Student t statistics estimated as a constant and trend in the regression of this test, we accept the hypothesis that they are not different to zero at a 5% significance level.

A value of –0.63 for statistic $\hat{t}_\tau$ of the ADF test was estimated for the LPEE series; this is lower than the 5% critical threshold of –3.45. Here again, this estimate is clearly lower than the critical values proposed by McKinnon (1991). As a result, we accept the null hypothesis of the presence of a unit root in the LPEE series. In another regression of this test (of order 5), we support the hypothesis that neither the constant nor the trend are significantly different to zero at a 5% significance level.

A significance level of 5% at –3.4049 (for order 6) was estimated to take into account the influence on the critical values of the ADF tests (Cheung and Lai, 1995). It was then verified that in absolute value the statistics $\hat{t}_\tau$ of the two variables, –0.50 and –0.63, were lower for this level of significance. In conclusion, the two variables are at least integrated of order 1, I(1).

The null hypothesis of the presence of two integrated unit roots of order 2, I(2), was then verified for each variable. The second differences (∇^2) for the series are shown in the lower part of Table 1. The criterion for the determination of the minimum lag order ($\hat{p}$ AIC+2, 10) selected an order $\hat{p} = 6$. This made it possible to overcome the problem of residual autocorrelation. Thus, for the LNI series, the statistic $\hat{t}_{ws,\tau}$ of the WS test was –4.40, which was distinctly higher than the 5% significance level of –3.36, with asymptotic probability (P value) of 0.0009. Statistic $\hat{t}_\tau$ of the ADF test was then estimated at –4.64, that is to say at a higher level than the 5% level of Fuller (1976) at –3.45 and of McKinnon (1991) at –4.08. The asymptotic probability (P value) of the ADF test was 0.001. The PP test gave negligible probability (1.9xd–07) for the null hypothesis. This led us to rejecting the null hypothesis of the presence of two unit roots for the LNI variable.

For the LPEE series, the statistic $\hat{t}_{ws,\tau}$ of the WS test was estimated to be –3.89, that is to say greater than its theoretical value at a 5% significance level, –3.36, with a P value of 0.006. The statistic $\hat{t}_{\tau}$ of the ADF test was found to be –5.65, which is distinctly greater than the 5% significance level at –3.45 on the one hand and McKinnon's estimates of –3.47 (5%), with a P value of 0.00001, on the other. We therefore rejected the null hypothesis of the presence of two unit roots in this series.

These tests failed to reject the hypothesis that each series can be represented by a difference stationary process (DS). Thus, the VAR model in which causality is analysed must be expressed in first differences. The stationarised series given by these tests are shown in Figures 3 and 4.

2. Spectral analysis

One of the greatest difficulties encountered in the analysis of quantitative phenomena stems from the fact that they are closely linked to time flow and one is obliged to observe them without being able to perform experiments involving the infinite repetition of the same circumstance. Any statistical approach thus includes the production of more or less artificial phenomena.

In the socioeconomic field, spectral analysis,[3] often presented as the most interesting of the procedures for detecting cycles, is no exception to the rule. Statistical series are often not long enough. They do not meet stationarity requirements and the elimination of the trend may affect the identification of the peaks.

Spectral analysis thus involves a regularity of movements that is not verified and which, in addition, is not essential in affirming that they exist. In fact, the spectral analysis method cannot truly prove or refute the existence of socioeconomic cycles. The non-pertinent nature of the method returns one to the more general problem of the non-neutrality of the method used with regard to the results obtained. Thus, cycles of different duration can be obtained that are not included in the basic data solely through the smoothing method used. Error in perspective is caused by the fact that the narrower the 'statistical window', the more chance there is of showing short cycles. Likewise, a broad 'statistical window' will accentuate long movements. From the theoretical point of view, one might consider that the problem could be solved by means of a statistical window covering the longest known movement in socioeconomic life, that is to say the Kondratieff (60 years). However, even in such a case,

[3] Spectral analysis is based on the theory of stochastic processes. The central hypothesis is that a time series consists of a large number of sinusoidal components with different frequencies (univariate spectral analysis). It makes it possible to divide a particular category of series into a set of oscillations with different frequencies and then to show the links between the components with the same frequency in the various series examined (cospectral or bivariate spectral analysis).

error in perspective would not be ruled out as long movements may be linked to even longer movements.

However that may be, in the present state of knowledge, spectral analysis is still the most suitable method for identifying or refining certain theoretical hypotheses. We are of course aware of the limits of such a procedure. In fact, from the epistemological point of view, we consider that although clearly raising questions of method is legitimate, it is pointless to complicate them indefinitely. It also seems obvious that socioeconomic series can never achieve the steadiness of the phenomena examined in natural science. In addition, it would be naive to hope to detect uniform chronology with symmetrical frequencies in socioeconomic life. It would seem today that the aim is more to propose a coherent interpretation with a strong heuristic value than to provide irrefutable proof of the existence of such patterns. Finally, our method was first of all determined by the search for the greatest possible objectivity in the observation of time series and then by the possibility of applying it to a large number of series. This two-fold requirement was dictated by the need to anticipate the criticism generally aimed at statistical studies on long movements in the economy.

Our in-depth analysis of the movements of German public expenditure on education and national income over a long period of time leads to a major observation: the spectral density functions of the stationary series showed a cycle of 7.9 years in economic growth and a cycle of 8.7 years in public expenditure on education. A 12.4 to 14.5 year movement was also observed in national income and one varying from 14.5 to 17.4 years in public expenditure on education.

This is an important result from the analytical point of view. It seems to confirm earlier work by Metz and Spree on growth of the German economy from 1820 to 1914 [Metz and Spree, 1981] and the analyses by Diebolt and Solomou [Diebolt, 1994; Solomou, 1987]. The latter studied GNP figures for Germany, the USA, France and the UK and observed only 'Kuznets' type *secondary secular movements* during the period 1850-1913.

The fact that the duration of the fluctuations was closer to Juglar and Kuznets cycles than to Kondratieff movements does not at all mean that the latter are statistical artefacts. On the contrary, the results oblige *supporters* of Kondratieff movements to revise their theories to match them with empirical results. However, the shifts in period are not surprising. They result from the confusion between the economic nature of the socio-economic cycles and the historical disturbances to which they may be subjected. Economic development is never as well-oiled as the movement of the stars, but marked with specifically historical traits that can affect or even change growth.

Figure 3

National Income in Germany, 1850-1937

First Differences

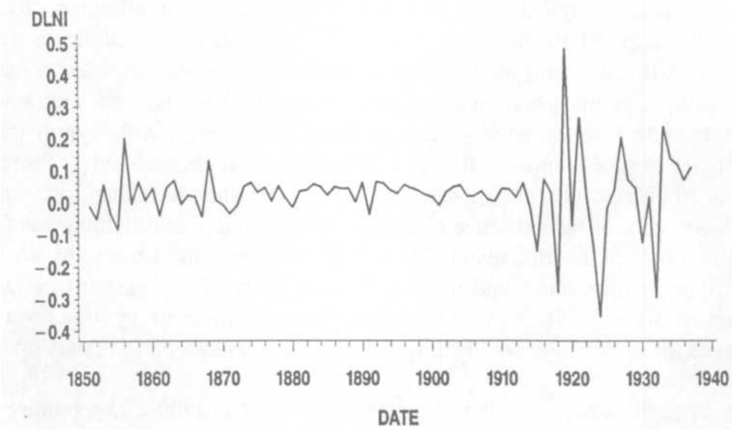
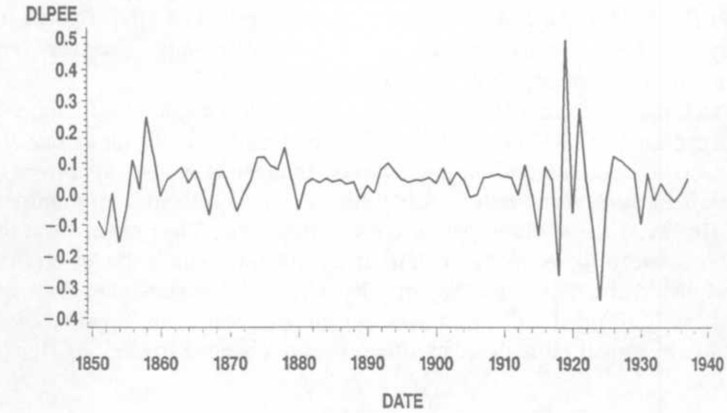


Figure 4

Public Expenditure on Education in Germany, 1850-1937

First Differences



3. Analysis of dynamic links

The stationarised series for national income and public expenditure on education reveal fluctuations of different types that may be synchronous or reversed (Figures 3 and 4). Cross-correlations, cospectra and finally causal relationships were used for closer analysis of the dynamic links between these variables. The results were used to seek a possible cointegration link. Cross-correlation analysis suggested a strong contemporaneous relationship between the series, ρ (DLNI, DLPEE) = 0.83 at time zero (no other significant cross-correlation was observed).

Cospectral analysis was first performed with no smoothing. The estimations of the *quadratic form* fluctuated continuously around zero, with a high coefficient of variation (854). Frequencies fluctuated between zero and π. The estimated mean was 0.00007. Student's t test (0.78) confirmed the null hypothesis for this mean with probability of 44%. In addition, the *quadratic form* displayed equivalent maximum and minimum values (0.0017 and –0.0016 at the frequencies 3.03 and 1.88). The cospectre itself displayed values that differed significantly from zero for most frequencies and which fluctuated considerably (the coefficient of variation was 123). Student's t test makes it possible to reject the null hypothesis of the mean (estimated to be 0.0013) with a probability of more than 0.99. The large peak with a frequency of $\omega = 0.72$ (period 8.7 years) should be noted here. The peaks with frequencies of 2.09 and 2.60 are smaller. It is deduced that the frequency components ω of the two variables studied are correlated and in phase with each other. This being stated, for reasons of scientific probity, we stress that we are aware that the estimators in a cospectral analysis should be smoothed by means of a spectral window. For the purposes of our research, we chose a (1 1 1 1 1) window and a triangular window (1 2 3 2 1). Smoothing confirmed the previous result obtained without smoothing.

A major result was found in causality tests. The Granger causality tests [Granger, 1969 ; Geweke, 1984] failed to reject the hypothesis that economic growth has a causal impact on education but made it possible to reject the reciprocal causal relationship. In other words, the causality tests make it possible to accept solely the causality hypothesis from the variable DLNI to the variable DLPEE, with 93% probability for the first lag. The F test gave the relatively modest value of 3.53. Following Durbin's h test, 0.6, the Breusch-Godfrey LM test (for lags 1 to 13) and the Ljung-Box Q(25) test, 28.6, we reject the hypothesis of residual autocorrelation.

Finally, the existence of cointegration links was tested. In a general manner, the cointegration concept can be described as follows: two series X_t and Y_t are said to be cointegrated of order (d, b), noted $(X_t, Y_t) \sim CI\ (d, b)$ if they are integrated of order d (their d differences are stationary) and if there is a linear combination of integrated series of order lower than d or integrated series $(d\text{-}b)$, where $b > 0$. Thus, two series are CI (1, 1) with a cointegration vector $(1, c)$ if

Figure 5

National Income in Germany, 1850-1937

Spectral Density (Frequency)

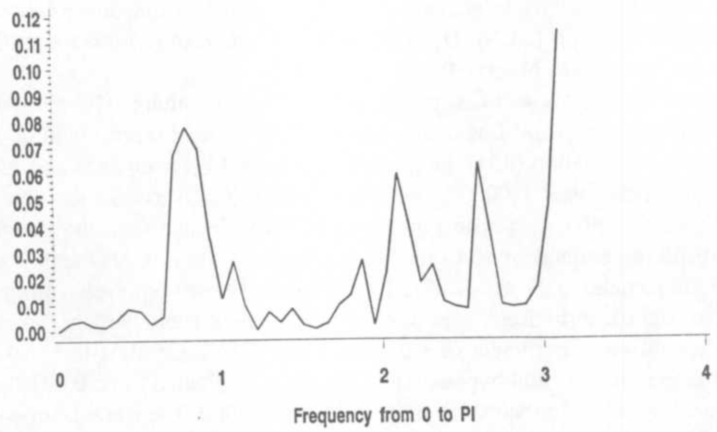


Figure 6

Public Expenditure on Education in Germany, 1850-1937

Spectral Density (Frequency)

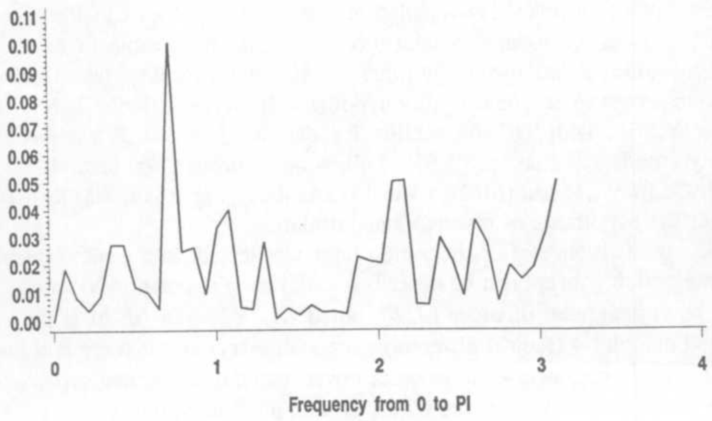


Figure 7

National Income in Germany, 1850-1937

Spectral Density (Period)

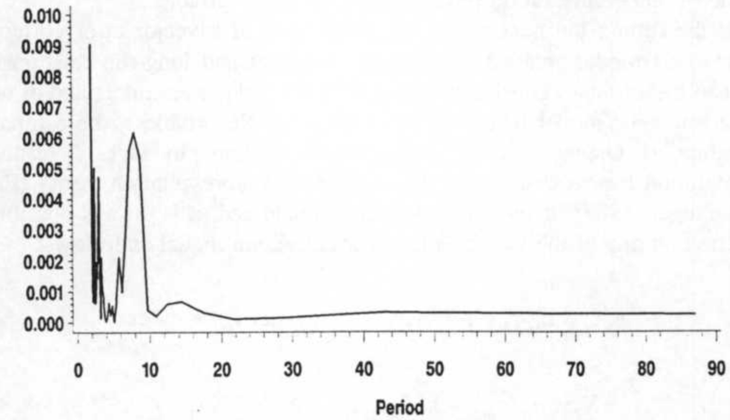
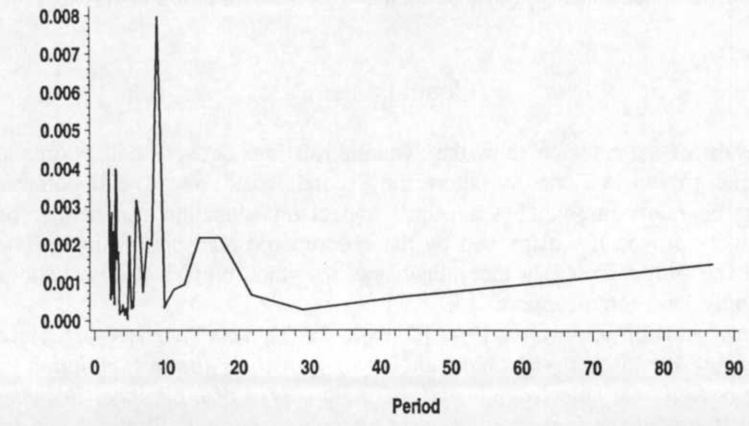


Figure 8

Public Expenditure on Education in Germany, 1850-1937

Spectral Density (Period)



they are stationary in prime differences but with a linear combination of their stationary level: $zt = X_t - cY_t$. This is an important property. Firstly, it implies the presence of relations of dependence and secondly it requires different formulation of the tests of causality. In this respect, neither the Engle-Granger test [Engle and Granger, 1987], the McKinnon test [Mc-Kinnon, 1991, 1994], nor the Johansen test [Johansen, 1988, 1991] rejected the hypothesis that there is at least one cointegration relation between these variables.

For the future, the next step is the estimation of a vector error correction model. This model enabled us to study short-run and long-run relationships between the variables considered. Indeed, as the series were integrated of order 1, the reference model is a VAR in differences. Nevertheless, the estimation procedure is changed for a cointegration relation. In fact, a result of cointegration is provided by Engle and Granger's representation theory [Engle and Granger, 1987]: if two variables are cointegrated, it is always possible to construct on one of the variables an error correction model as follows:

$$\Delta X_t = a_0\, z_{t-1} + \Sigma b_{1i} \Delta X_{t-i} + \Sigma b_{2i} \Delta X_{t-i} + e_t, \text{ or}$$

$$\Delta X_t = a'_0\, z_{t-1} + \Sigma b'_{1i} \Delta X_{t-i} + \Sigma b'_{2i} \Delta X_{t-i} + e'_t$$

where $z_{t-1} = X_{t-1} - cY_{t-1}$, and with e_t and e'_t stationary.

The significance of a_0 (respectively of a'_0) reveals Y to X causality (and X to Y respectively) because X (and respectively Y) adjusts following a shock that moves the variables away from their cointegration relation. A vector error correction model (VECM) rather than a VAR must then be used to test causality. In short, term z_{t-1} must be included in the estimation of VAR.

Conclusion

The results of the research show the dynamic relations between education and economic growth in Germany before the Second World War. The hypothesis that the economic growth has a causal impact on education can neither be definitively proved nor disproved by the econometric analysis of single time series. The proposition is a theoretical one, to which empirical observations might only lend some support.

A new situation arose after 1945 Diebolt, 1997]. This suggests that, since the Second World War *»the main engine of growth is the accumulation of human capital - of knowledge - and the main source of differences in living standards among nations is differences in human capital. Physical capital* *accumulation plays an essential but decidedly subsidiary role.«* [Lucas, 1993, p. 270]

Bibliography

Akaike, H.: »Maximum Likelihood Estimation of Gaussian Autoregressive Moving Average Models«, in: *Biometrika,* 60, 1973, pp. 255-265.

Aubin, H.; Zorn, W. (Eds.): *Handbuch der deutschen Wirtschafts- und Sozialgeschichte,* Vol. 2: *Das 19. und 20. Jahrhundert,* Klett-Cotta, Stuttgart, 1976.

Barro, R.J.; Sala-i-Martin, X.: *Economic Growth,* McGraw-Hill, Inc., New York, 1995.

Becker, G.S.: *Human Capital. A Theoretical and Empirical Analysis with a Special Reference to Education,* Columbia University Press, New York, 1964.

Box, G.; Jenkins, G.: *Time Series Analysis. Forecasting and Control,* Holden Day, London, 1970. Bresson, G.; Pirotte, A.: *Econometrie des séries temporelles. Théorie et Applications,* Presses Universitaires de France, Paris, 1995. Cheung, Y.-W.; Lai, K.: »Lag Order and Critical Values of the Augmented Dickey-Fuller Test«, in: *Journal of Business and Economic Statistics,* 13, 1995, pp. 277-280.

Dickey, D.; Fuller, W.: »Distribution of the Estimators for Autoregressive Time Series with a Unit Root«, in: *Journal of the American Statistical Association,* 74, 1979, pp. 427-431.

Dickey, D.; Fuller, W.: »Likelihood Ratio Statistic for Autoregressive Time Series with a Unit Root«, in: *Econometrica,* 49, 1981, pp. 1057-1072.

Diebolt, C: L'*évolution de longue période du système éducatif allemand.* Vol. 1: *Education, croissance et cycles longs.* Vol. 2: L'*histoire quantitative de l'éducation,* (PhD diss. in Economics, University of Montpellier I, 15 February 1994.

Diebolt, C: »New Findings in Comparative Education: Towards a Theory of Systemic Regulation. The Case of France and Germany in the 19th and 20th Centuries«, in: Schriewer, J. in collaboration with M.J. Halverson (Eds.): *Discourse Formation in Comparative Education,* P. Lang, Frankfurt am Main, 1997.

Engle, R.; Granger, C: »Co-Integration, and Error Correction: Representation, Estimation, and Testing«, in: *Econometrica,* 55, 1987, pp. 251-276.

Engle, R.; Granger, C: *Long-Run Economic Relationships. Readings in Cointegration,* Oxford University Press, Oxford, 1991.

Engle, R.; Yoo, B.: »Forecasting and Testing in Co-Integrated Systems«, in: *Journal of Econometrics,* 35, 1987, pp. 143-159.

Freeman, C. (Ed.): *Long Wave Theory,* E. Elgar, Cheltenham, 1996. Fuller, W.: *Introduction to Statistical Time Series,* Wiley, New York, 1976. Geweke, J.: »Inference and Causality in Economic Time Series Models (Chapter 19)«, in: Grilliches, Z.; Intrilligator, M. (Eds.): *Handbooks of Econometrics,* Vol. 2, North Holland, Amsterdam, 1984, pp. 1101-1144. Godfrey, L.: »Testing against General Autoregressive and Moving Average Error Models when the Regressors Include Lagged Dependent Variables«, in: *Econometrica,* 46, 1978, pp. 1293-1310. Godfrey, L.: »Testing the Adequacy of a Time Series Model«, in: *Biometrika,* 66, 1979, pp. 67-72. Gourieroux, C; Monfort, A.: *Séries temporelles et modèles dynamiques,* 2ème édition, Economica, Paris, 1995. Granger, C: Investigating Causal Relationship between Econometric Methods and Cross-Spectral Methods«, in: *Econometrica,* 37, 1969, pp. 424-439.

Hamilton, J.: *Time Series Analysis,* Princeton University Press, New Jersey, 1994.

Hoffmann, W.G. *et al.: Das Wachstum der deutschen Wirtschaft seit der Mitte des 19. Jahrhunderts,* Springer-Verlag, Berlin, 1965. Johansen, S.: »Statistical Analysis of Cointegration Vectors«, in: *Journal of Economic Dynamics and Control,* 12, 1988, pp. 231-254. Johansen, S.: »Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models«, in: *Econometrica, 59,* 1991, pp. 1551-1580.

Juglar, C: *Des crises commerciales et de leur retour périodique en France, en Angleterre et aux Etats-Unis,* 2ème édition, Guillaumin et Cie, Paris, 1889.

Kleinknecht, A.; Mandel, E.; Wallerstein, I. (Eds.): *New Findings in Long-Wave Research,* St. Martin's Press, New York, 1992.

Kondratieff, N.D.: *Les grands cycles de la conjoncture,* Economica, Paris, 1992.

Kunst, R.; Neusser, K.: »Cointegration in a Macro-economic System«, in: *Journal of Applied Econometrics,* 5, 1990, pp. 351-365. Kuznets, S.: *Secular Movements in Production and Prices. Their Nature and their Bearing upon Cyclical Fluctuations,* Houghton Mifflin Co., Boston, 1930.

Kuznets, S.: »The Pattern of U.S. Economic Growth«, in: Phelps, E.S. (Ed.): *The Goal of Economic Growth. An Introduction to a Current Issue of Public Policy,* W.W. Norton & Co., New York, 1969, pp. 191-208.

Ljung, G.; Box, G.: »On a Measure of Lack of Fit in Time Series Models«, in: *Biometrika,* 65, 1978, pp. 297-303.

Lucas, R.E.: »On the Mechanics of Economic Development«, in: *Journal of Monetary Economics,* 22, 1988, pp. 3-42.

Lucas, R.E.: »Making a Miracle«, in: *Econometrica,* 61, 1993, pp. 251-272.

Lütkepohl, H.: *Introduction to Multiple Time Series Analysis,* Springer-Verlag, Berlin, 1991.

McKinnon, J.: »Critical Values for Co-Integration Tests«, in: Engle, R.F.; Granger, C.W.J. (Eds.): *Long-Run Economic Relationships,* Oxford University Press, Oxford, 1991, pp. 267-276.

McKinnon, J.: »Approximate Asymptotic Distribution Functions for Unit-Root and Cointegration Tests«, in: *Journal of Business and Economic Statistics,* 12, 1994, pp. 167-176.

Metz, R.; Spree, R.: »Kuznets-Zyklen im Wachstum der deutschen Wirtschaft während des 19. und frühen 20. Jahrhunderts«, in: Petzina, D.; Roon, G. van (Eds.): *Konjunktur, Krise, Gesellschaft. Wirtschaftliche Wechsellagen und soziale Entwicklung im 19. und 20. Jahrhundert,* Klett-Cotta, Stuttgart, 1981, pp. 343-376.

Nelson, C; Plosser, C: »Trend and Random Walks in Macroeconomic Time Series«, in: *Journal of Monetary Economics,* 10, 1982, pp. 139-162.

Pantula, S.; Gonzalez-Farias, G.; Fuller, W.: »A Comparison of Unit-Root Test Criteria«, in: *Journal of Business and Economic Statistics,* 12, 1994, pp. 449-459.

Priestley, M.B.: *Spectral Analysis and Time Series, 2* vol., Academic Press, New York, 1981.

Phillips, P.: »Time Series Regression with a Unit Root«, in: *Econometrica,* 55, 1987, pp. 277-301

Phillips, P.; Perron, P.: »Testing for a Unit Root in Time Series Regression«, in: *Biometrika,* 75, 1988, pp. 335-346.

Slutzky, E.: »The Summation of Random Causes as the Source of Cyclic Processes«, in: *Econometrica,* 5, 1937, pp. 105-146.

Solomou, S.: *Phases of Economic Growth, 1850-1973. Kondratieff Waves and Kuznets Swings,* Cambridge University Press, Cambridge, 1987.

Spree, R. unter Mitarbeit von M. Tybus: *Wachstumstrends und Konjunkturzyklen in der deutschen Wirtschaft von 1820 bis 1913. Quantitativer Rahmen für eine Konjunkturgeschichte des 19. Jahrhunderts,* Vandenhoeck & Ruprecht, Göttingen, 1978.

Thome, H.: »Trends, Cycles, and Co-Integration. Some Issues in Modelling Long-Term Development in Time Series Analysis«, in: *Historical Social Research,* 21, 1996, pp. 3-23.

Tortella, G. (Ed.): *Education and Economic Development since the Industrial Revolution,* Generalität Valenciana, Valencia, 1990.